Testing the regional restructuring hypothesis in western Germany

Herman J. Bierens
Department of Economics, Pennsylvania State University,
608 Kern Graduate Building, University Park, PA 16802, USA;
e-mail: hbierens@psu.edu

Thomas Kontuly
Department of Geography, University of Utah,
Salt Lake City, UT 84112, USA; e-mail: kontuly@geog.utah.edu

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Abstract

In this paper we test the importance of the regional restructuring hypothesis by investigating whether the internal migration flows in Western Germany can be explained by employment changes, on the basis of data of out-migration of employed workers and employment in 75 regions over the period 1982-1997 for each year separately. Starting from a conditional probability model for the unobserved individual migration decisions, we derive mathematically a dynamic version of the Poisson gravity model for the total out-migration per region per year. Estimating this model for each year separately, we find partial confirmation of the regional restructuring hypothesis.

1 Introduction

A change occurred in interregional population redistribution patterns in developed countries during the 1970s. This change was referred to as the ”migration turnaround”, the ”turnaround”, or the ”rural renaissance” in the USA, and was labeled ”counterurbanization” in Europe (Beale 1977; Berry 1976; Champion 1989; Vining and Kontuly 1978). Before 1970 regional (net) internal migration tendencies redistributed national populations in developed nations from smaller to larger sized urban or metropolitan areas. This traditional tendency was also characterized as an urbanization pattern. The migration turnaround became evident when internal redistribution patterns favored small- and medium-sized places at the expense of large urban areas. This phenomenon was also called counterurbanization.

Then during the 1980s a number of developed countries experienced a ”migration turn-back-around”, by reverting to traditional patterns of urbanization (Champion 1992; Cochrane and Vining 1988; Richter 1985; Rogerson and Plane 1985). Other nations showed either continuing or a new pattern of counterurbanization. While researchers were attempting to comprehend the long-term implications of the turnaround and turn-back-around in the latter part of the 1980s, a fundamental transformation occurred in European migration patterns as a result of the collapse of Communist political systems in Europe.

Western Germany, as the old Federal Republic of Germany, closely followed the changing patterns of interregional movement described above. Between 1970 and 1984 a trend toward greater interregional deconcentration (or counterurbanization) occurred in western Germany, measured in terms of a net shift down the urban size hierarchy from large to small-sized urban regions (Kontuly and Vogelsang 1988). This shift occurred within the context of a general north-to-south redistribution of population that reflected the economic and industrial transformations taking place in the country. A shift back toward spatial concentration or an urbanization tendency, with net migration occurring up the urban hierarchy, started in 1985 and continued through 1988 (Kontuly et al 1997). As a result of these changes it was not clear if regional tendencies were returning to traditional patterns of concentration, were a temporary postponement of deconcentration tendencies or represented alternating cycles of spatial concentration and deconcentration.

In 1989 a major change occurred in German migration patterns due to the collapse of the former Soviet Union. Then the political reunification of

Germany took place in October 1990. During the years 1989 to 1994, a total of 1.1 million Germans relocated from the eastern to the western part of the country (Kontuly et al 1997). Then in the post-unification period of the 1990s, an interregional deconcentration of the population was found once again for the western part of Germany (Kemper 2004).

Deindustrialization started in the 1970s in western Germany, and this economic decline caused by the demise of the coal, iron and steel industries was regionally specific and concentrated in the Rhine-Ruhr area and the Saarland. Since 1970, Germany transformed itself from a manufacturing to a service based economy. By 1999 the percent of the labor force employed in manufacturing declined to 33.4 while employment in the service sector increased to 63.8%, with only 2.8% employed in agriculture (CIA 2007). This process of economic restructuring impacted large urban areas and rural areas alike (Gans and Kemper 2002).

The literature suggests numerous explanations for either a slowing urbanization pattern or the existence of a counterurbanization tendency in European countries during the 1970s, 1980s, and 1990s. These explanations can be organized into five groups, that is, regional restructuring, the deconcentration perspective, government policy, housing costs, and period effects. Regional restructuring refers to the regional deconcentration of manufacturing and service employment, and can be thought of as the "people follow jobs" explanation. The deconcentration perspective focuses on the location choices of individuals and refers to a change in the preferences of the working-age population in favor of locations in rural, low density, environmentally pleasant areas; this is the "jobs follow people" explanation. Period effects include business cycle fluctuations, regional boom and bust experiences, and changing socio-demographic compositions. The government policy explanation includes planned deconcentration initiatives to redistribute jobs and people to rural / peripheral areas, and includes increases in public service employment such as, child care, education for the young, and other services. Housing costs reasons include both the high cost and the availability of housing (Kontuly 1998; Kontuly and Dearden 1998, Kemper 2004; Audirac 2005).

Several researchers believe that the first three explanations provide the basis for understanding counterurbanization and slowing urbanization in the developed world (Fielding 1989; Frey 1987, 1988, 1989, 1993; Hugo and Smailes 1985; Kontuly and Vogelsang 1988; Morrill 1978; Moseley 1984; Plane and Rogerson 1991; Audirac 2002; Kemper 2004). Kontuly and Vogelsang (1988) contend that government policy must be added to these three major explanations for the case of western Germany.

The regional restructuring and the deconcentration perspectives explain long-term tendencies in the human settlement patterns of developed countries. Regional restructuring explanations stress the global changes in industrial production and the changing functions of metropolitan areas as a result of the new spatial division of labor (Elliott 1997; Frey and Speare 1988; Noyelle and Stanback 1984). This perspective contended that slowing urbanization reflects the selective depopulation of metropolitan areas dependent upon traditional manufacturing and forecasted new urbanization tendencies for large metropolitan areas that were able to function as advanced private service centers, the locations of corporate headquarters, banks, and other financial institutions, and for metropolitan areas that became the location of knowledge-based industries. Smaller metropolitan areas and non-metropolitan regions relying on routine industrial production would almost certainly experience decline (Elliott 1997). This perspective draws from a wide variety of theoretical traditions (Audirac 2005), such as world systems theory (Timberlake 1985), world city formation (Sassen 1991), the work on flexible production by Storper and Scott (1992) and Scott (1988), and writings on the information and networked society (Castells 1989, 1996). Regional restructuring appeared in the 1980s to explain American and European deindustrialization (Audirac 2005).

The deconcentration perspective argues that a change in residential preferences of the working-age population or in the ability to act on such preferences occurred in favor of residences in rural or small-town environments and against large cities. This perspective was developed mainly by demographers, human ecologists, regional scientists, and urban geographers to explain the "migration turnaround" evident in the 1970s (Audirac 2005). This perspective was found in the literature to be an important explanation in four out of the fourteen countries reviewed. These countries were France, Western Germany, Norway and Switzerland (Kontuly 1998).

Period explanations, which account for short-term change, attributed the 1970s migration turnaround to the unique economic and demographic circumstances of the 1970s, and implied that when these economic and demographic shocks subsided, more traditional urbanization tendencies would re-emerge. The changing socio-demographic compositions explanation included the influence of relative age cohort sizes such as the proportion of the population in the Baby Boom generation and the proportion of the population that is elderly. The emergence of large birth cohorts during the 1950s and early 1960s impacted labor markets and the spatial structure of young adult migration flows became more geographically focused. The aging of the population has meant a rapid increase in the numbers of economically inactive people and a growing pool of potential migrants. Retirement enables those with sufficient resources to realize long-held desires for more realized life-styles in environmentally attractive, less congested surroundings.

Regional restructuring is occurring as the proportion of tertiary and quaternary employment increases relative to secondary employment. Also, the decline in primary employment has almost run its course, so there is a reduction in the stock of potential out-migrants living in rural areas. Older industrial countries have been going through a process of deindustrialization that has had a strong negative impact on larger cities, especially on their central areas. Fielding (1989) expands on this explanation by emphasizing the evolution of a new spatial division of labor. He argues that the growth of companies into multi-plant, multi-product, and multi-national enterprises was accompanied by disinvestment in high-wage areas in favor of low-wage reserves of labor within national territories. The deconcentration of jobs was followed by population migration.

Regional restructuring is one of the major factors suggested in the literature as an explanation for changing urbanization patterns in European countries during the 1970s and 1980s. Regional restructuring or economic structural change was cited in the literature as an important explanation in eleven out of fourteen countries surveyed. These countries were Belgium, Great Britain, Denmark, Finland, France, Western Germany, Italy, the Netherlands, Norway, Sweden, and Switzerland (Kontuly 1998). Renewed interest in internal migration processes in developed countries and the reasons behind the changes in these processes was stimulated by the discovery of a "new migration turnaround" in the USA during the 1990s (Fuguitt and Beale 1996; Johnson 1999; Long and Nucci 1997). A similar "new" turnaround was evident in western Germany during the period 1992- 1996 (Kemper 1999), and Kemper (2004) argues that the contrasting internal migration trends in eastern and western Germany during the 1990s are consistent with the expectations of the regional restructuring hypothesis.

Elliott (1997) evaluated the importance of the regional restructuring versus the deconcentration perspective in a study of population migration in U. S. metropolitan statistical areas between 1965 and 1990. His analysis found little confirmation for either of these two dominant perspectives. Instead, Elliott's (1997) "results show that 'concentrated dispersion' or deconcentration, has in fact increased over recent decades in a manner consistent with theories of systemic maturation" (p. 37). In contrast to both the deconcentration and regional restructuring perspectives, systemic maturation contends that individual metropolitan areas co-evolve as both the cause and consequence of long-standing migratory processes that bind the national settlement system. As proponents of systemic maturation, Geyer and Kontuly (1993) argue that migration processes are best conceptualized in terms of three successive stages of "differential urbanization". Kontuly and Dearden (2003) found that the differential urbanization or systemic maturation model accurately characterized urban population development in Western Germany between 1939 and 2010. According to them (2003, p.64), progression through the stages of the model did not occur "in lock step fashion, but was evident as a consistent tendency in a general direction". Elliott also argues that his "findings are important for ongoing efforts to theorize recent population redistribution trends because they imply that contemporary economic change is not the only force driving metropolitan-based migration. Historic and enduring processes of population deconcentration also continue to play a key role" (Elliott 1997 p. 39).

The literature discussing possible explanations for either slowing urbanization or counterurbanization in western Germany provided support for the regional restructuring, the deconcentration, and the systemic maturation perspectives (Kontuly 1998: Kontuly and Dearden 2003). This paper will focus on the regional restructuring hypothesis and attempt to assess quantitatively its importance by determining the degree to which internal migration flows in Western Germany can be related to employment change differences. The association between yearly changes in the Western Germany migration system and changes in employment will be evaluated over a long period, that is, during the 1980s and 1990s.

The plan of the paper is as follows. After a brief discussion of the available data, our starting point in the next section is a conditional probability model of the individual migration decisions, which relates the probability to migrate from region j to region i to the past migration flow from region j to region i, and the difference of the employment change rates in regions i and j. Upon aggregation this procedure gives rise to a multinomial model for the migration flows, which can be approximated by a dynamic version of the Poisson gravity model proposed by Flowerdew and Aitken (1982). Since we do not observe these migration flows, the Poisson gravity model is further aggregated into a dynamic Poisson model for the out-migration in each region. In Section 3 we present the maximum likelihood estimation results for each time period separately, and these results are discussed and interpreted in relation to the regional restructuring hypothesis. In particular, we find only partial confirmation of the regional restructuring hypothesis, in the sense that the key parameter, which measures the effect of employment opportunities elsewhere on out-migration, flip-flops signs over time. In section 4 we summarize our empirical findings, and we suggest avenues for further research.

In the Appendix we address some econometric issues. First, we compute 90% confidence intervals of the parameter estimates by bootstrapping, because due to the small sample size the asymptotic t-values may not be accurate enough. The bootstrap results indicate that we were too pessimistic about this problem. Second, the validity of the model specification is tested by the Hausman (1978) test. Third, we provide web links to the software we have used.

2 Modeling migration at the regional level

2.1 The data

For $j = 1, ..., N = 75$ regions in West Germany and years $t = 1982, ..., 1997$ we observe:

$$\begin{aligned} y_{j,t}^{-} &= \text{out-migration of employed workers,} \\ y_{j,t}^{+} &= \text{in-migration of employed workers,} \\ Z_{j,t} &= \text{total employment.} \end{aligned}$$

Neither migration nor employment are age-restricted. If the migration data involved would concern genuine **internal** migration, then for each year the net migration should add up to zero:

$$\sum_{j=1}^{N} y_{j,t}^{-} - \sum_{j=1}^{N} y_{j,t}^{+} = 0. \tag{1}$$

However, this is not the case in our data set, and therefore the migration data contain in-migration from, and out-migration to, areas outside the 75 regions under review. This problem is likely not too serious a problem for the out-migration, but it is for the in-migration, especially after the German unification. Therefore, only the out-migration data can be assumed to represent internal migration. Whether this assumption is correct or not does not matter too much for the model, as will become clear.

2.2 Methodology

Let $y_{i|j}(t)$ be the total number of employed workers (exclusive of family members) migrating from region j to region i in year t. Then the out-migration of employed workers (exclusive family members) from region j is

$$y_{j,t}^{-} = \sum_{i=1}^{N} I(i \neq j) y_{i|j}(t) \tag{2}$$

where I is the indicator function[1], and if the add-up restriction (1) holds the in-migration in region i would be given by

$$y_{i,t}^{+} = \sum_{j=1}^{N} I(i \neq j) y_{i|j}(t). \tag{3}$$

Note that $Z_{j,t} = Z_{j,t-1} + y_{j,t}^{+} - y_{j,t}^{-}$, ignoring the number of employed workers who died in period t.

We do not observe the internal migration flows $y_{i|j}(t)$, but they play a important role in the formation of the model. To formula a logically consistent model for the migration flows $y_{i|j}(t)$ we start with a plausible probability model for the unobserved individual decision to migrate from region i to region j in year t. Aggregating the individual decisions to the region level then yields a conditional multinomial probability model for the migration flows $y_{i|j}(t)$. Using the Poisson approximation of the multinomial distribution, and aggregrating the migration flows further to total in- and out-migration per region, then yield a dynamic version of the Poisson gravity model proposed by Flowerdew and Aitken (1982). Thus, the core of our modeling approach is the individual probability model involved. Given the latter model, the dynamic Poisson gravity models for $y_{j,t}^{-}$ and $y_{i,t}^{+}$ are derived mathematically.

[1] $I(true) = 1$, $I(False) = 0$.

2.3 A probability model for individual migration decisions

Since moving takes time, we assume that the **decision to migrate takes effect in the next year**. Consequently, the total number of employed workers (exclusive of family members) in region j that decide stay in region j is

$$y_{j|j}(t) = Z_{j,t-1} - y^{-}_{j,t}. \tag{4}$$

The model we propose is a model for the conditional probability $p_{i|j}(t)$, given the past, that in period t a representative **employed** worker in region j will migrate to region $i \neq j$.

First, it is plausible that $p_{i|j}(t)$ is positively related to the **fraction** of workers who migrated from region i to region j in period $t-1$, i.e., we assume that $p_{i|j}(t)$ depends on

$$\frac{y_{i|j}(t-1)}{Z_{j,t-2}}. \tag{5}$$

Since we assume that the decision to migrate in period $t-1$ is made in period $t-2$, the ratio (5) is the fraction of workers in region j in period $t-2$ who have decided to migrate to region i. Moreover, note that by (2) and (4),

$$\sum_{i=1}^{N} \frac{y_{i|j}(t-1)}{Z_{j,t-2}} = 1. \tag{6}$$

The fraction (5) measures the attractiveness of region i as migration destination for workers in region j. One may consider this as a "keeping up with the Joneses" effect.

Moreover, it will be assumed that the decision to migrate from region j to region i is also based on a variable

$$X_{i,j,t-1} \in [0,1], \tag{7}$$

representing the employment opportunities in region i relative to the employment opportunities in region j in year $t-1$. This variable will be specified later.

This suggests the following specification of the conditional probability $p_{i|j}(t)$ that in period $t \geq 3$ an employed worker in region j migrates to region $i \neq j$:

$$p_{i|j}(t) = \alpha_0 \frac{y_{i|j}(t-1)}{Z_{j,t-2}} + \beta_0 \frac{X_{i,j,t-1}}{N} + \gamma_0 \frac{1}{N}. \tag{8}$$

with γ_0/N playing the role of "intercept".

In order for (8) to be a probability distribution, we need to require that for $i \neq j$, $p_{i|j}(t) \in [0,1]$, and $\sum_{i=1}^{N} I(i \neq j) p_{i|j}(t) < 1$. Hence, assuming that $\alpha_0 > 0$, $\beta_0 > 0$, and $\min_{i\neq j} X_{i,j,t-1} = 0$, $\max_{i\neq j} X_{i,j,t-1} = 1$, we must have

$$\begin{aligned}
&\alpha_0 \min_{i\neq j} \frac{y_{i|j}(t-1)}{Z_{j,t-2}} + \gamma_0 \frac{1}{N} \geq 0, \\
&\alpha_0 \max_{i\neq j} \frac{y_{i|j}(t-1)}{Z_{j,t-2}} + (\beta_0 + \gamma_0) \frac{1}{N} \leq 1, \\
&\alpha_0 \frac{y_j^-(t-1)}{Z_{j,t-2}} + \beta_0 \frac{\sum_{i=1}^{N} I(i \neq j) X_{i,j,t-1}}{N} + \gamma_0 \frac{N-1}{N} < 1.
\end{aligned} \tag{9}$$

Consequently, it is possible that $\gamma_0 < 0$.

Note that the restrictions (9) involve random variables, and therefore the coefficients α_0, β_0 and γ_0 are in fact random variables themselves.

2.4 Poisson approximation of the multinomial distribution

As is well-known, the general form of the probability function of the multinomial distribution is

$$q(y_1, .., y_{k-1}) = \frac{n!}{y_1!....y_{k-1}!\left(n - \sum_{j=1}^{k-1} y_j\right)} \left(\prod_{j=1}^{k-1} p_j^{y_j}\right) \left(1 - \sum_{j=1}^{k-1} p_j\right)^{n - \sum_{j=1}^{k-1} y_j}. \tag{10}$$

Define $\lambda_j = np_j$ and suppose that $p_j \downarrow 0$ as $n \to \infty$ such that the λ_j's remain constant. Denote $z = \sum_{j=1}^{k-1} y_j$. Then similar to the Poisson approximation of the Binomial distribution (see for example Bierens 2004, pp.9-10) it can be shown that for fixed $y_1, .., y_{k-1}$ and $n \to \infty$,

$$q(y_1, .., y_{k-1}) \to \prod_{j=1}^{k-1} \left[\exp(-\lambda_j) \left(\lambda_j^{y_j}/y_j!\right)\right]. \tag{11}$$

Hence, for large n and small p_j's the multinomial distribution is approximately equal to the distribution of $k-1$ **independent** Poisson(λ_j) variates, where $\lambda_j = np_j$.

2.5 The model

Assuming that the probabilities $p_{i|j}(t)$ are small and $Z_{j,t-1}$ is large, it follows now that for fixed j the random variables $y_{i|j}(t)$are approximately independent Poisson$\left(Z_{j,t-1}p_{i|j}(t)\right)$ distributed, conditional on the past. Moreover, since for $j = 1,..,N$ the random variables $y_{i|j}(t)$ are due to the decisions of employed workers in different regions, we may without loss of generality assume that, conditional on the past, the $y_{i|j}(t)$'s are independent Poisson$\left(Z_{j,t-1}p_{i|j}(t)\right)$ distributed for all unequal i and j. Therefore, adopting the specification (8) we have approximately:

Conjecture 1 *For $i, j = 1, ..., N$, $i \neq j$ and each t the migration flows $y_{i|j}(t)$ from region j to region i are independent Poisson$(\eta_{ijt}(\alpha_0, \beta_0, \gamma_0))$ distributed, conditional on the past, where*

$$\eta_{ijt}(\alpha, \beta, \gamma) \quad = \quad \left(\alpha_0 \frac{y_{i|j}(t-1)}{Z_{j,t-2}} + \beta_0 \frac{X_{i,j,t-1}}{N} + \gamma_0 \frac{1}{N}\right) Z_{j,t-1} \qquad (12)$$

Note that this is a special case of the Poisson gravity model proposed by Flowerdew and Aitken (1982). See also, for example, Okoruwa et al. (1988, 1994) for applications.

Moreover, note that the migration flows $y_{i|j}(t)$ are not independent over time, because they depend on past values. Furthermore, it is unlikely that the parameters are constant over time. But even if they are we cannot estimate them jointly for all t because the time dimension is too small to use laws of large numbers and central limit theorems for dependent random variables. Thus, the parameters will be estimated **for each time period** t **separately.**

If the migration flows $y_{i|j}(t)$ and $y_{i|j}(t-1)$ were observable, the parameters could be estimated by maximum likelihood on the basis of the result in Conjecture 1. However, we only observe total in and out-migration per region. In order to get around this problem, we will use the well-known fact that sums of independent Poisson variates are Poisson distributed themselves, i.e., it follows from Conjecture 1 that conditional on the past, $y^+_{i,t}$ is Poisson $\left(\sum_{j=1}^N I(j \neq i)\eta_{ijt}(\alpha_0, \beta_0, \gamma_0)\right)$ distributed, and $y^-_{j,t}$ is Poisson

$\left(\sum_{i=1}^{N} I(i \neq j)\eta_{ijt}(\alpha_0, \beta_0, \gamma_0)\right)$ distributed:

$$\begin{aligned} y_{i,t}^{+} &\sim \text{Poisson}\left(\alpha_0 \sum_{j=1}^{N} I(j \neq i) y_{i|j}(t-1) \frac{Z_{j,t-1}}{Z_{j,t-2}}\right. \\ &\left. +\beta_0 \frac{1}{N} \sum_{j=1}^{N} I(j \neq i) X_{i,j,t-1} Z_{j,t-1} + \gamma_0 \frac{1}{N} \sum_{j=1}^{N} I(j \neq i) Z_{j,t-1}\right) \end{aligned} \tag{13}$$

and

$$\begin{aligned} y_{j,t}^{-} &\sim \text{Poisson}\left[\left(\alpha_0 \frac{y_{j,t-1}^{-}}{Z_{j,t-2}} + \beta_0 \left(\frac{1}{N} \sum_{i=1}^{N} X_{i,j,t-1} - \frac{X_{j,j,t-1}}{N}\right)\right.\right. \\ &\left.\left. +\gamma_0 \left(1 - \frac{1}{N}\right)\right) Z_{j,t-1}\right] \end{aligned} \tag{14}$$

Since we do not observe the $y_{i|j}(t-1)$'s, only the latter result is usable.

2.6 Employment opportunities

Since the main purpose of this study is to investigate to what extent employment has an impact on internal migration, the parameter β_0 is the key parameter. However, in order to identify β_0 we need to specify $X_{i,j,t-1}$ such that there is enough variation in $(1/N)\sum_{i=1}^{N} X_{i,j,t-1}$ for $j = 1, ...N$. Asymptotically, i.e., for $N \to \infty$, (14) becomes

$$y_{j,t}^{-} \sim \text{Poisson}\left[\left(\alpha_0 \frac{y_{j,t-1}^{-}}{Z_{j,t-2}} + \beta_0 \lim_{N\to\infty} \frac{1}{N} \sum_{i=1}^{N} X_{i,j,t-1} + \gamma_0\right) Z_{j,t-1}\right], \tag{15}$$

so that without enough variation in $(1/N)\sum_{i=1}^{N} X_{i,j,t-1}$ we will run into a near-multicollinearity problem. A specification of $X_{i,j,t-1}$ that will just avoid this problem is, for example,

$$X_{i,j,t-1} = I\left(Z_{i,t-1} > Z_{j,t-1}\right) \tag{16}$$

or

$$X_{i,j,t-1} = I\left(\Delta \ln Z_{i,t-1} > \Delta \ln Z_{j,t-1}\right), \tag{17}$$

where Δ is the difference operator, so that $(1/N)\sum_{i=1}^{N} X_{i,j,t-1}$ is then the fraction of regions with larger employment, or larger employment growth, respectively, in year $t-1$ than in region j. Clearly, in either case

$$\bigcup_{j=1}^{N} \left\{\frac{1}{N} \sum_{i=1}^{N} X_{i,j,t-1}\right\} = \left\{0, \frac{1}{N}, \frac{2}{N}, ..., \frac{N-1}{N}\right\}. \tag{18}$$

If we would choose the specification (16), it will be difficult to distinguish employment effects from scale effects. Therefore, our preferred choice is specification (17).

Note that $X_{j,j,t-1} = 0$. Moreover, without loss of generality we may redefine $\gamma_0 \left(1 - \frac{1}{N}\right)$ in (14) as γ_0. Thus, the model we will estimate is:

$$y_{j,t}^{-} \sim \text{Poisson}\left[\left(\alpha_0 \frac{y_{j,t-1}^{-}}{Z_{j,t-2}} + \beta_0 F_{N,j,t-1} + \gamma_0\right) Z_{j,t-1}\right] \tag{19}$$

where

$$F_{N,j,t-1} = \frac{1}{N} \sum_{i=1}^{N} I\left(\Delta \ln Z_{i,t-1} > \Delta \ln Z_{j,t-1}\right). \tag{20}$$

2.7 Discussion

Now let us reconsider the issue whether the out-migration $y_{j,t}^{-}$ represents genuine internal out-migration. Suppose it does not. Then we should have taken an additional region, "outside Western Germany", into account, as region $i = 0$ for example, with "employment" $Z_{0,t}$. Consequently, we then should have used in model (19)

$$F_{N,j,t-1}^{*} = \frac{1}{N+1} \sum_{i=0}^{N+1} I\left(\Delta \ln Z_{i,t-1} > \Delta \ln Z_{j,t-1}\right) \tag{21}$$

as the variable representing employment opportunities elsewhere rather than $F_{N,j,t-1}$. However, the error involved is small: $\left|F_{N,j,t-1}^{*} - F_{N,j,t-1}\right| \leq 2/\left(N+1\right)$ ≈ 0.0263, so that ignoring the region "outside Western Germany" will not affect the model substantially.

Model (19), together with the assumption $\beta_0 > 0$, imply that people have a preference for faster growing regions as migration destination. However, it is also conceivable that some people have an opposite preference, because fast growing regions are also characterized by traffic congestion, expensive housing, and other problems. Taking these opposite preferences into account, we may rewrite model (19) as

$$y_{j,t}^{-} \sim \text{Poisson}\left[\left(\alpha_0 \frac{y_{j,t-1}^{-}}{Z_{j,t-2}} + \beta_0^{+} F_{N,j,t-1} + \beta_0^{-}\left(1 - F_{N,j,t-1}\right) + \gamma_0\right) Z_{j,t-1}\right], \tag{22}$$

where $\beta_0^+ \geq 0$, $\beta_0^- \geq 0$. Denoting $\beta_0^* = \beta_0^+ - \beta_0^-$, $\gamma_0^* = \beta_0^- + \gamma_0$, we can write (22) as

$$y_{j,t}^- \sim \text{Poisson}\left[\left(\alpha_0 \frac{y_{j,t-1}^-}{Z_{j,t-2}} + \beta_0^* F_{N,j,t-1} + \gamma_0^*\right) Z_{j,t-1}\right]. \tag{23}$$

Therefore, a negative sign of β_0 is possible.

3 Empirical results

The parameters of model (19) have been estimated by maximum likelihood, without imposing any restrictions. The estimation results together with the corresponding asymptotic t-values are presented in Table 1 below, where $\widehat{\alpha}$, $\widehat{\beta}$ and $\widehat{\gamma}$ are the maximum likelihood estimators of α_0, β_0 and γ_0, respectively. The asymptotic t-values of α_0 are twofold: one for testing the null hypothesis $\alpha_0 = 0$, and the other one for testing the null hypothesis $\alpha_0 = 1$.

As we suspected, the parameters are not constant over time. Moreover, recall that both β_0 and γ_0 may take negative values, and therefore the same applies to $\widehat{\beta}$ and $\widehat{\gamma}$.

Table 1: *Maximum Likelihood Estimation results*

Year	$\widehat{\alpha}$	$\widehat{t}(\alpha_0 = 0)$	$\widehat{t}(\alpha_0 = 1)$	$\widehat{\beta}$	$\widehat{t}(\beta_0 = 0)$	$\widehat{\gamma}$	$\widehat{t}(\gamma_0 = 0)$
1984	0.8682957	325.684	−49.400	0.0081801	35.324	0.0002114	0.747
1985	0.9698154	297.275	−9.252	0.0018213	7.053	0.0041130	15.607
1986	0.9297883	309.616	−23.380	0.0004035	1.709	0.0034148	12.067
1987	1.0047547	327.848	1.551	−0.0010500	−4.660	−0.0013386	−4.609
1988	1.0913222	377.611	31.599	0.0082802	35.930	−0.0109299	−36.661
1989	1.3236419	550.239	134.538	−0.0060175	−25.925	−0.0135048	−52.398
1990	1.0276992	523.472	14.109	0.0089300	35.426	−0.0028939	−12.917
1991	0.5640496	370.518	−286.371	0.0042057	18.404	0.0311582	150.815
1992	0.8481089	377.274	−67.568	−0.0001668	−0.818	0.0157178	72.574
1993	0.7365150	274.508	−98.204	0.0094785	41.556	0.0187003	66.957
1994	1.0679468	379.901	24.171	−0.0059622	−26.721	0.0033641	12.334
1995	1.0504782	424.561	20.401	0.0050018	20.925	−0.0064779	−20.596
1996	0.8484628	354.112	−63.245	0.0010321	4.557	0.0128886	43.622
1997	0.9550732	349.050	−16.419	0.0026204	11.426	0.0044223	14.393

The results for the $\widehat{\beta}$'s are at first sight puzzling. A positive sign would indicate that a majority of the migrants favor the fast growing employment regions over slow growing areas as migration destinations, and a negative sign indicates the other way around. Thus, the positive $\widehat{\beta}$'s seem to corroborate the regional restructuring hypothesis or the "people follow jobs" explanation of internal migration discussed in the Introduction. Negative $\widehat{\beta}$ coefficients cannot be interpreted as confirmation of the deconcentration hypothesis. The fluctuations of the coefficients are not in phase with variations in the business cycle nor with periods of prevalent urbanization or counterurbanization. The $\widehat{\beta}$ coefficients indicate whether the migration balance of employed people is positively or negatively correlated with changes in employment. Since the growth of employment is higher in the rural regions of western Germany during the period analyzed in this paper (Kemper 2004), the negative $\widehat{\beta}$ coefficients indicate a migration tend toward urbanized areas in spite of below average employment growth. We tested Kemper's statement with the data evaluated in this paper, and found that employment change was relatively higher in rural regions compared to large urban regions in every year 1984 to 1997. Migration during years in which the $\widehat{\beta}$ coefficients are negative could be the result of a majority of migrants moving for non-employment related reasons, such as movement to attend University or to secure an apprenticeship position with a firm. Such movers would most likely be in the 18 to 24 year old age group. If the bulk of the migrants, in a particular year, were found within this age group then the negative $\widehat{\beta}$ coefficients could reflect this type of non-employment related movement. Since human migration is highly age selective, it is not unreasonably that total migration would be dominated by the moves of the young. Migrants within the 25 to 29 year old age group moved primarily for employment related reasons (Kemper 1999) and responded to changes in the West German labor market during the 1980s, so we would expect to find positive coefficients for this age group. Movers in the 30 to 49 year old family group could move for employment or non-employment related reasons. Disaggregating migration by the different age groups (18-24, 25-29, 30-49, 50-64 and 65 and over) could solve the puzzle of the signs of the $\widehat{\beta}$'s flip-flopping over time. Separate analyzes would need to be performed for each age group.

In order to explain the sign changes of the $\widehat{\beta}$'s, suppose that in year $t-1$, $100 \times q_{t-1}\%$ of the people in our sample favor living in a high-growth region: Their probability of migrating from region j to region i in year t is:

$$p_{i|j}(t|t-1) \quad = \quad \alpha_* \frac{y_{i|j}(t-1)}{Z_{j,t-2}} + \beta_* \frac{I\left(\Delta \ln Z_{i,t-1} > \Delta \ln Z_{j,t-1}\right)}{N}$$

$$+\gamma_* \frac{1}{N}. \tag{24}$$

Moreover, suppose that the other$100 \times (1 - q_{t-1})\%$ favor living in a low-growth region: Their probability of migrating from region j to region i in year t is:

$$\begin{aligned} p_{i|j}(t|t-1) &= \alpha_* \frac{y_{i|j}(t-1)}{Z_{j,t-2}} + \beta_* \frac{1 - I\left(\Delta \ln Z_{i,t-1} > \Delta \ln Z_{j,t-1}\right)}{N} \\ &\quad + \gamma_* \frac{1}{N}. \end{aligned} \tag{25}$$

For the sake of clarity of the argument the parameters $\alpha_* > 0, \beta_* > 0$ and γ_* are assumed to be the same in both cases. Moreover, assume that the fraction q_{t-1} applies to all regions. Then if we randomly select an employed worker the probability that he or she will migrate from region j to region i is

$$\begin{aligned} p_{i|j}(t|t-1) &= \alpha_* \frac{y_{i|j}(t-1)}{Z_{j,t-2}} + \left(2q_{t-1} - 1\right)\beta_* \frac{I\left(\Delta \ln Z_{i,t-1} > \Delta \ln Z_{j,t-1}\right)}{N} \\ &\quad + \left((1 - q_{t-1})\beta_* + \gamma_*\right) \frac{1}{N} \end{aligned} \tag{26}$$

Now suppose that $\beta_* = 0.1$. If $q_{t-1} = 0.51$, then $\beta_0 = \left(2q_{t-1} - 1\right)\beta_* = 0.002$, and if $q_{t-1} = 0.49$ then $\beta_0 = \left(2q_{t-1} - 1\right)\beta_* = -0.002$. This example shows that even small changes in preferences may result in significant sign changes of β_0. Therefore, our results in Table 1 do not exclusively support one of the explanations for internal migration reviewed in the introduction. Instead, the sign changes of the $\widehat{\beta}$'s indicate that each of these explanations may apply to different subsets of employed workers.

The variation of the $\widehat{\alpha}$'s over time may be explained similar to the variation of the $\widehat{\beta}$'s.

4 Concluding remarks

The literature suggests three possible explanations for changing regional urbanization patterns in Western Germany during the 1980s and 1990s, that is, the regional restructuring perspective, the deconcentration perspective, and the systemic maturation perspective. In this paper we focus on the regional restructuring perspective and investigate the extent that inter-regional migration in Western Germany can be explained by economic factors. Yearly interregional labor force out-migration, for the 75 regions of Western Germany, is related to changes in regional employment over the period 1982 to 1997. Aggregating the conditional probabilities to migrate from one area to another yields a conditional Poisson model for out-migration from each of the 75 regions, with conditional expectation depending on previous out-migration relative to employment and a variable representing employment opportunities elsewhere. This model is estimated for each year separately, and its validity is tested by the Hausman test. In view of the small sample size (75), bootstrap 90% confidence intervals of the parameters were constructed to verify the asymptotic t-values, and the bootstrap results indicate that the asymptotic t values are reliable. Also bootstrap p-values were constructed for the Hausman test, and the test results do not provide evidence that the Poisson model is misspecified.

We find that the β coefficients of the employment opportunity variable switch signs over time. Coefficients with a positive sign indicate that a majority of migrants favor regions with rapidly growing employment, while coefficients with a negative sign indicate the opposite. In our study, positive signs on the β coefficients of the employment change variable were found for ten of the fourteen years analyzed, and significant negative signs in three of the fourteen years. A negative coefficient could result from migration toward large-size, densely populated urban regions in spite of below average employment growth or migration for non-economic reasons. For example, migration to attend University or to secure an apprenticeship position with a firm. We find only partial confirmation for the regional restructuring hypothesis and conclude that economic factors cannot exclusively explain inter-regional migration patterns in Western Germany. Elliott (1997) arrived at a similar conclusion in an analysis of migration among US metropolitan areas between 1965 and 1990.

It is suggested that future research undertake an age-disaggregated evaluation of internal migration in Western Germany, in order to understand better the reasons for the negative signs found in this analysis. For example, employment availability provides the main attraction to migrants between the ages of 25 and 29, because this age group moves out of university areas and regions with high unemployment rates for the young (Kontuly and Vogelsang 1988). Evaluating yearly associations between internal migration and employment change for this age group is then expected to generate positive

β coefficients.[2]

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[2] However, incorporating this age-disaggregated evaluation in the current paper would increase the size of the paper beyond the page limit of *Environment & Planning A,* and would cause unduly time delay. The second author plans to undertake this further research together with Ph.D. students on the basis of the methodology in the current paper.

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Appendix

Bootstrap quantiles

Since the sample size ($N = 75$) is small, the asymptotic t-values in Table 1 may not be suitable for testing the significance of the parameters involved. Therefore, we have bootstrapped the 5 and 95 % quantiles of the maximum likelihood estimates as follows. Given $Z_{j,t-1}, Z_{j,t-2}, F_{N,j,t-1}$, and the maximum likelihood estimates $\widehat{\alpha}$, $\widehat{\beta}$ and $\widehat{\gamma}$, we have generated for each t 1000 replications of $y_{j,t}^{-}$, $j = 1, ..., 75$, by drawing each $y_{j,t}^{-}$ independently from the Poisson($\widehat{\lambda}_j(t)$) distribution, where

$$\widehat{\lambda}_j(t) = \left(\widehat{\alpha}\frac{y_{j,t-1}^{-}}{Z_{j,t-2}} + \widehat{\beta}F_{N,j,t-1} + \widehat{\gamma}\right) Z_{j,t-1}. \tag{27}$$

Actually, the Poisson random drawings are generated using the normal approximation of the Poisson distribution: if Y is Poisson(λ) distributed and λ is large, then $(Y - \lambda)/\sqrt{\lambda}$ is approximately $N(0, 1)$ distributed. Therefore, the $y_{j,t}^{-}$'s are generated as $y_{j,t}^{-} = \max\left(0, \widetilde{\lambda}_j(t) + u_j\sqrt{\widetilde{\lambda}_j(t)}\right)$, rounded off to integers, where the u_j's are independent random drawings from the standard normal distribution.

For each replication $r = 1, ..., 1000$, the maximum likelihood estimators $\widetilde{\alpha}_r$, $\widetilde{\beta}_r$ and $\widetilde{\gamma}_r$ are computed, and after completion the estimates $\widetilde{\alpha}_r$, $\widetilde{\beta}_r$ and $\widetilde{\gamma}_r$ are sorted in increasing order. Then the 5% quantile of each parameter corresponds to $r = 50$, and the 95% quantile corresponds to $r = 950$. The two quantiles form the lower and upper bounds of the 90% confidence interval of each parameter.

The results, presented in Table 2, indicate that using asymptotic t-values for testing parameter significance yield the same conclusions as using the bootstrapped 90% confidence intervals. Thus, we were too pessimistic about the validity of the asymptotic results.

Table 2: *Bootstrap quantiles*

Year	$\widehat{\alpha}$	$\widehat{\alpha}$	$\widehat{\beta}$	$\widehat{\beta}$	$\widehat{\gamma}$	$\widehat{\gamma}$
	5%	95%	5%	95%	5%	95%
1984	0.8639591	0.8726062	0.0077612	0.0085491	−0.0002696	0.0006526
1985	0.9644676	0.9750794	0.0013555	0.0022383	0.0036932	0.0045300
1986	0.9251099	0.9349646	0.0000283	0.0007903	0.0029597	0.0038525
1987	0.9999429	1.0093696	−0.0014128	−0.0006778	−0.0018003	−0.0008677
1988	1.0863572	1.0963706	0.0079084	0.0086522	−0.0114396	−0.0104429
1989	1.3193744	1.3279010	−0.0063973	−0.0056374	−0.0139466	−0.0130659
1990	1.0245302	1.0309348	0.0085311	0.0093655	−0.0032754	−0.0025465
1991	0.5615442	0.5665677	0.0038339	0.0045942	0.0308046	0.0314791
1992	0.8445624	0.8518055	−0.0005255	0.0001740	0.0153681	0.0160808
1993	0.7324421	0.7408718	0.0091056	0.0098324	0.0182503	0.0191264
1994	1.0635456	1.0727262	−0.0063543	−0.0055976	0.0029064	0.0038084
1995	1.0465234	1.0547731	0.0046013	0.0053857	−0.0070160	−0.0059697
1996	0.8444956	0.8525165	0.0006567	0.0013937	0.0124307	0.0133739
1997	0.9504109	0.9592600	0.0022351	0.0029724	0.0039361	0.0049695

The Hausman test

The maximum likelihood estimators have been computed in two steps. First, initial values $\widetilde{\theta}_{OLS} = (\widetilde{\alpha}_{OLS}, \widetilde{\beta}_{OLS}, \widetilde{\gamma}_{OLS})'$ have been computed by regressing $y_{j,t}^{-}$ on $\frac{y_{j,t-1}^{-}}{Z_{j,t-2}}Z_{j,t-1}$, $F_{N,j,t-1}Z_{j,t-1}$ and $Z_{j,t-1}$. Then the vector $\widehat{\theta} = (\widehat{\alpha}, \widehat{\beta}, \widehat{\gamma})'$ of maximum likelihood estimators is computed by the Newton iteration, starting from $\widetilde{\theta}_{OLS}$. A by-product of this procedure is the Hausman test.

Under the null hypothesis that the model is correctly specified, i.e., the hypothesis that (19) is true, both $\widetilde{\theta}_{OLS}$ and $\widehat{\theta}$ are asymptotically normally distributed around the true parameter vector $\theta_0 = (\alpha_0, \beta_0, \gamma_0)'$, i.e.,

$$\sqrt{N}\left(\widetilde{\theta}_{OLS} - \theta_0\right) \to N_3\left[0, V_{OLS}\right], \ \sqrt{N}\left(\widehat{\theta} - \theta_0\right) \to N_3\left[0, V_{ML}\right], \qquad (28)$$

in distribution as $N \to \infty$, and since the maximum likelihood estimator is asymptotically efficient, $V_{OLS} - V_{ML}$ is positive semi-definite. Hausman

(1978) has shown that then $\sqrt{N}\left(\widetilde{\theta}_{OLS}-\widehat{\theta}\right) \rightarrow N_3\left[0, V_{OLS}-V_{ML}\right]$. On the other hand, if the hypothesis (19) is false, then one may expect that

$$H_1\text{: } p\lim_{N\rightarrow\infty}\sqrt{N}\left\|\widetilde{\theta}_{OLS}-\widehat{\theta}\right\| = \infty. \tag{29}$$

Given consistent estimates $\widehat{V}_{OLS}$ and $\widehat{V}_{ML}$ of V_{OLS} and V_{ML}, respectively, and assuming that $V_{OLS}-V_{ML}$ is positive definite, it follows that under the null hypothesis (19),

$$\widehat{H} = N\left(\widetilde{\theta}_{OLS}-\widehat{\theta}\right)\left(\widehat{V}_{OLS}-\widehat{V}_{ML}\right)^{-1}\left(\widetilde{\theta}_{OLS}-\widehat{\theta}\right) \rightarrow \chi_3^2 \tag{30}$$

in distribution as $N \rightarrow \infty$, whereas under the alternative hypothesis (29), $p\lim_{N\rightarrow\infty}\widehat{H} = \infty$. The statistic $\widehat{H}$ is the test statistic of the Hausman test.

In order to check the model specification (19) we have conducted the Hausman test, with $\widehat{V}_{OLS}$ White's (1980) heteroskedasticity-consistent covariance matrix estimator, and

$$\widehat{V}_{ML} = \left(\frac{-1}{N}\left.\frac{\partial^2 \ln \mathcal{L}(\theta)}{\partial\theta\partial\theta'}\right|_{\theta=\widehat{\theta}}\right)^{-1} \tag{31}$$

where $\mathcal{L}(\theta)$ is the likelihood function. However, one should bear in mind that:

1. $V_{OLS}-V_{ML}$ may have zero eigenvalues, and therefore $\widehat{V}_{OLS}-\widehat{V}_{ML}$ may have zero or even negative eigenvalues;

2. The sample size $N = 75$ is small, so that the asymptotic χ_3^2 distribution may not be close;

3. The power of the test may be poor, not only because N is small, but also because there is no guarantee that (29) holds if (19) is false.

Problem 1 can be solved by replacing $\widehat{V}_{OLS}-\widehat{V}_{ML}$ by

$$\widehat{\Sigma} = \widehat{Q}\begin{pmatrix} \widehat{\Lambda}_+ & O \\ O & O \end{pmatrix}\widehat{Q}' = \widehat{Q}_+\widehat{\Lambda}_+\widehat{Q}_+' \tag{32}$$

where $\widehat{\Lambda}_+$ is the diagonal matrix of positive eigenvalues of $\widehat{V}_{OLS}-\widehat{V}_{ML}$, and $\widehat{Q} = \left(\widehat{Q}_+, \widehat{Q}_-\right)$is the orthogonal matrix of corresponding eigenvectors. Then under the null hypothesis (19),

$$\widehat{H} = N\left(\widetilde{\theta}_{OLS}-\widehat{\theta}\right)\widehat{Q}_+\widehat{\Lambda}_+^{-1}\widehat{Q}_+'\left(\widetilde{\theta}_{OLS}-\widehat{\theta}\right) \rightarrow \chi_q^2 \tag{33}$$

in distribution, where q is the rank of $V_{OLS} - V_{ML}$.

Problem 2 can be solved by bootstrapping the null distribution. This is what we have done in Table 3 below. For each of the $r = 1, ..., 1000$ bootstrap replications we have computed the Hausman test statistic $\widetilde{H}_r$ in the form (33). The bootstrap p-value is then the frequency of the event $\widetilde{H}_r > \widehat{H}$. The asymptotic p-values are based on the χ^2_3 distribution.

Table 3: *Hausman test*

Year	Test statistic $\widehat{H}$	Asymptotic p-value	Bootstrap p-value
1984	3.767	0.28769	0.27000
1985	1.349	0.71748	0.46800
1986	0.831	0.84210	0.58400
1987	1.811	0.61248	0.41200
1988	0.763	0.85838	0.55000
1989	1.789	0.61740	0.47900
1990	2.658	0.44741	0.27600
1991	1.284	0.73303	0.42500
1992	0.858	0.83548	0.51300
1993	0.340	0.95241	0.67100
1994	2.014	0.56946	0.42900
1995	4.485	0.21368	0.23100
1996	1.018	0.79699	0.46200
1997	0.657	0.88331	0.52100

As we see from Table 3, the Hausman test does not reject the null hypothesis at any conventional significance level. Note that all but one of the bootstrap p-values are substantially smaller than the asymptotic p-values, indicating that the asymptotic χ^2 null distribution is a poor approximation of the small sample null distribution.

Software

The estimation of the model in this paper has been conducted via a special plug-in module of the free econometric software package *EasyReg International* (shortly EasyReg) developed by the first author. This software package can be downloaded from web page http://econ.la.psu.edu/~hbierens/EASYREG.HTM. How to estimate the model via EasyReg and to prepare EasyReg for this task is demonstrated on web page http://econ.la.psu.edu/

~hbierens/EasyRegTours/POISSONG.HTM. The latter web page also contains the data involved.